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Competition and Concentration in the EU Banking Industry*

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I. Introduction

Competition and concentration in the European financial system are major issues these days. Indeed, the establishment of Economic and Monetary Union (EMU) is generally expected to change the financial landscape in the European Union (EU) dramatically (Molyneux and Gardner, 1997; Vander Vennet, 1997). It is argued, for example, that the creation of large and transparent euro capital markets will enhance competition in the European banking industry and stimulate disintermediation and securitisation. Eventually, this would lead to further consolidation and rationalisation in the European banking sectors. In this respect, it is also argued that EMU will uncover different starting positions of banking sectors in individual countries. Vives (1991) suggests that some countries have relatively concentrated banking sectors, whereas others are faced with excess capacity in their banking sectors. As a consequence, he asserts that significantly different competitive conditions exist in banking markets across EU countries. Moreover, EMU will also increase the pressure for further harmonisation of regulation across EU countries, so that the incentives and opportunities for regulatory arbitrage will diminish considerably. In this paper, we shall empirically test the validity of these qualitative notions about competition and concentration in the European banking industry, in particular the structure-conduct-performance paradigm, which states that increased concentration fosters collusion and anti-competitive practices.

There are also more general reasons why market conditions in the banking industry deserve particular attention. First, the orthodox view holds that increasing concentration in some segments of the banking

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market may eventually result in an undesirable exercise of market power by banks. The argument is generally supported by an appeal to concentration ratios, e.g. the relative size of the five largest banks. Second, the soundness and stability of the financial sector may in various ways be influenced by the degree of competition and concentration. Nowadays, safeguarding the health of the financial system is one of the key objectives of bank supervisors.

Somewhat surprisingly, empirical investigations of competitive conditions and concentration in EU banking industries are scarce. The bulk of recent empirical studies on competition in the banking sector pertains to the United States (Shaffer, 1982), Canada (Nathan and Neave, 1989) and Japan (Lloyd-Williams et al., 1991). This paper tries to fill this gap and extends previous studies on competition and concentration in European banking in several respects. We examine the overall market conditions in which banks of all fifteen current EU Member States operate over the period 1989 - 96. The underlying model is based on an amended version of the Panzar-Rosse methodology (Rosse and Panzar, 1977; Panzar and Rosse, 1982, 1987). We explicitly account for possible gradual changes in the market structure by introducing a logistic time-curve model into the original Panzar-Rosse framework, which allows a pooled cross-section and time-series analysis. This way, the analysis also yields information about the effects of the liberalisation and deregulation of international capital markets on the competitiveness of European banking markets, an element that has been neglected in existing studies up to now. The modified framework is applied to the EU as a whole as well as to separate national banking sectors. Another novelty is that we attach different weights to the banks included in our sample; shares of individual banks in the assets of all banks are taken as weights. As will be explained, a weight-based model is preferable from an economic point of view. Subsequently, we link the issues of competition and concentration. Earlier empirical studies on the European banking industry often focus on just one of these factors. We use the assessment of the competitive nature of banking markets with the modified Panzar-Rosse approach to establish the relationship between competitive conduct and degree of concentration. By doing so, we implicitly test the validity of either the structure-conduct-performance or the contestability/efficiency paradigm in European banking.

The structure of this paper is as follows. Section II. briefly discusses methodological and institutional issues regarding competition and concentration in European banking sectors. Here, the available empirical evidence for European countries is also summarised. Section III. presents the Panzar-Rosse model used in our econometric examination to assess competitive conditions in the banking markets. Section IV. contains the estimation results of this model. Section V. investigates the link between competition and concentration in the banking sector. The final section presents our conclusions.

II. Methodological and institutional considerations

The determination of the degree of competition and concentration in the banking industry and the relationship between these concepts pose both analytical and methodological questions. Basically, the issue at stake is how the "natural" market structure of the banking sector can be characterised. This largely depends on the existence of scale or scope efficiencies or, more precisely, multiproduct cost subadditivity in banking (Baumol, 1977). Numerous empirical studies have been devoted to this topic (e.g. Berger et al., 1993; Molyneux et al., 1996). Most studies reach the conclusion that the market conditions prevailing in banking sectors can best be characterised as naturally oligopolistic. This means that – in the long term – there is only room for a few viable banks. In this context, a highly concentrated banking sector is the logical outcome of market forces.

Competition in banking is often analysed by reference to income and cost structures of banking sectors (Berger et al., 1993). Generally speaking, two basic methods can be distinguished to determine the competitive nature of the banking industry empirically. The first approach is due to Bresnahan (1982) and comes down to a simultaneous estimation of a market demand or supply function and a price setting equation using industry aggregate figures. From this exercise, a parameter indexing the oligopoly solution concept (λ) is identified by standard econometric methods. The comparative statics of equilibrium, as price and quantity are moved by exogenous variables, reveal the degree of market power. If λ equals zero, perfect competition exists. If $\lambda = 1$, there is a perfect cartel. Intermediate $\lambda's$ correspond to other oligopoly solution concepts.

The second approach, used in this study, is developed by Panzar and Rosse (1987) and requires firm-specific data. They have constructed a so-called H statistic to make a quantitative assessment of the competitive nature of banking markets and the market power of banks. The H statistic is calculated from reduced-form revenue equations and measures the sum of elasticities of total revenue of the bank with respect to the bank's input prices. Panzar and Rosse show that this statistic can reflect the structure and conduct of the market in which the bank operates. This interpretation underscores that the competitive environment faced by the bank is not necessarily identical to the competitive situation prevailing in the country where the bank is located. This remark holds particularly for large universal banks with sizeable foreign activities. These internationally active banks are obviously confronted by other competitive forces than small regional banks. For countries with relatively closed banking systems, the H statistic is thus more indicative of the competitive situation in the domestic banking market.

Concerning the value of H, Panzar and Rosse assert that H is zero or negative when the competitive structure is a monopoly, a perfectly colluding oligopoly, or a conjectural variations short-run oligopoly. Under these conditions, an increase in input prices will increase marginal costs, reduce equilibrium output and subsequently reduce total revenues. Under perfect competition, the H statistic is unity. In this case, an increase in input prices raises both marginal and average costs without – under certain conditions – altering the optimal output of any individual firm. Exit of some firms increases the demand faced by each of the remaining firms, thereby leading to an increase in prices and total revenues by the same amount as the rise in costs. Monopolistic competition models are *a priori* most plausible for characterising the interaction between banks. The monopolistic competition model recognises the existence of product differentiation and is consistent with the observation that banks tend to differ with respect to various product quality variables and advertising, although their core business is fairly homogeneous. Panzar and Rosse prove that, under monopolistic competition, H is unity or less. H is a decreasing function of the perceived demand elasticity, so H increases with the competitiveness of the banking industry. The testable hypotheses are: the banking industry is characterised by monopoly for $H \leq 0$, monopolistic competition for $0 < H < 1$ and perfect competition for $H = 1$. Panzar and Rosse (henceforth PR) treat banks as single-product firms, using deposits and other funding costs as inputs to produce merely loans and other interest earning assets. This is the so-called intermediation approach to bank modelling, which emphasises the financial intermediation role of banks. The approach cannot be applied to separate segments of the banking market. Therefore, the observed H is an index of the overall competitive conditions in the whole banking market.

As noted earlier, there have been few empirical studies assessing competitive conditions in European banking markets so far. Suominen (1994) and Swank (1995) apply the approach of Bresnahan (1982) to estimate demand and supply functions for Finnish and Dutch banks, respectively. Suominen finds evidence of imperfect competition, whereas Swank concludes that in the latter part of the sample period 1957 - 1990, Dutch mortgage and savings deposits markets were significantly more oligopolistic than in Cournot equilibrium. Molyneux et al. (1994) use the Panzar-Rosse (PR) methodology in a sample of German, United Kingdom, French, Italian, and Spanish banks for each year of the period 1986 to 1989. On average, their results point to monopolistic competition in Germany, France, Spain and the UK and to monopoly in Italy. Actually, the results of Molyneux et al. are rather unstable over the successive years. For example, the market structure faced by banks in the UK shifted from monopoly to almost perfect competition and back. In the case of German banks, the H statistic switched signs between 1986 and 1989 (–0.0363 in 1986; 0.4697 in 1989). It is highly unlikely, of course, that competitive conditions change that drastically in a few years' time. Vesala (1995) applies a similar model to the Finnish banking industry and finds monopolistic competition for 1985 - 88 and 1991 - 92, and perfect competition for 1989 - 90.

Regarding the question of concentration and capacity in the European banking market, more studies are available (Baltensperger and Dermine, 1990; Caminal et al., 1990; among others). In these studies, the share of the assets of the three, five or ten largest banks in total bank assets, branch numbers and numbers of banks are mostly used to proxy concentration and capacity levels in banking markets. Without exception, these articles reveal a striking feature of European banking markets (Molyneux and Forbes, 1995). In almost every country, a handful of large banks tends to emerge over time, whether through government encouragement or through market mechanisms.

Having identified the competitive structure and level of concentration in banking markets, the question arises what the relationship between these concepts is. From the theoretical literature, it appears that the nature of this link is ambiguous. In certain circumstances, a highly concentrated banking sector can impair competition. The maintained assumption is then that concentration translates into greater market power, thus leading to collusive behaviour and excess profits for the financial institutions (Gual and Neven, 1992). This is in line with the traditional structure-conduct-performance (SCP) paradigm. The paradigm states that increased concentration fosters collusion and anti-competitive practices.[1]

This line of reasoning is challenged by two other theoretical strands. The *contestability theory* stresses that a concentrated banking industry can behave competitively if the hurdles to be surmounted by new entrants to the market are low (Baumol, 1982). This theory asserts that the threat of potential entry forces banks with large market shares to price their products competitively under certain conditions. In a perfectly contestable market, entry is absolutely free, exit is completely costless and the demands for industry outputs are highly price-elastic. A contestable market has no entry barriers, neither economic nor legal. Costless exit implies that when a firm is planning to enter a new market, it expects to recover fixed costs (for instance, through sale of assets) if it later decides to withdraw. These features and highly price-elastic demands imply that a contestable market is effectively competitive even if it has only a small number of active firms. If the proponents of the contestability theory are correct, widely expressed concerns about the domination of a country's financial system by some type of financial intermediaries may be valid only to the extent that financial markets are not contestable (Nathan and Neave, 1989). On the other hand, the *efficiency hypothesis* states that if a bank enjoys a higher degree of efficiency than its competitors, it can adopt two different strategies. The first option is to maximise profits by maintaining the present levels of prices and firm size. The second alternative is to maximise profits by reducing prices and expanding firm size. If the bank chooses the second option, the most efficient banks will gain market share and bank efficiency will be the driving force behind the process of market concentration. Hence, both the contestability and the efficiency hypothesis assume

1 Here, we acknowledge that some segments of the markets are presumably somewhat more competitive than others. For example, it may be difficult for banks to acquire and exercise substantial market power in the market for large corporate loans, because many – foreign – banks compete for that business. If large firms consider the price of bank credits too high, they could easily decide to enter the international capital markets directly for obtaining financial means; that is, without using the intermediary services of banks. Empirically, this notion is buttressed by *Neven* and *Röller* (1995) for seven European countries (i.e. Belgium, Denmark, France, Germany, Spain, the Netherlands and the United Kingdom). They have estimated a structural model over the period 1981 - 1989 for wholesale and retail markets in which the degree of collusion, the price elasticity of demand and the interest margin are included. The hypothesis of the existence of identical market power in both markets is rejected. The degree of collusion in the retail segment appears to be higher than *vis-à-vis* clients in wholesale markets. In this article we consider *average* competitive conditions.

that the overall competitive environment faced by banks does not necessarily depend on the degree of market concentration.

For European countries, one of the crucial assumptions underlying the latter hypotheses, i.e. free entry and exit, was definitely not satisfied until the late 1980s. It is a well-known fact that government intervention in most of the banking systems in Europe was fairly extensive up to the mid-1980s. Banks were also legally restricted in their activities both within their national markets and across borders. Some countries applied capital restrictions or erected high obstacles to the establishment of branches by foreign banks (Bisignano, 1992). This tight regulatory environment was relaxed considerably with the adoption of the Second Banking Co-ordination Directive of 1992 as part of the single European market project.[2] In the first place, this directive boosted the deregulation and liberalisation of international capital flows. This process was accompanied by the formulation and implementation of other policy initiatives, such as the lifting of restrictions on interest payments on deposits and the development of a harmonised framework for supervision of the European banks. In the beginning of 1993, all formal restrictions regarding the provision of financial services across the European Union were removed. Banks which are licensed anywhere in the Union are given a "single banking licence", which allows them to service the entire European market, either by setting up branches in other countries or by offering products across national borders. According to the European Commission (1997), these developments have spurred competition and concentration in European banking sectors. There are also indications that it has improved the overall efficiency of European banking markets (Groeneveld, 1999).

This paper captures these considerations in various ways. We first determine the competitive structure of the entire European banking industry using an extended version of the PR method. To take into account the possibility that the variables under consideration are in some way or another influenced by the institutional changes sketched above, we have incorporated a logistic time-curve model in the original PR specification. Subsequently, we examine the competitive behaviour of banks in individual EU countries. The estimated H statistics are used to investigate the relationship between competition and concentration. This enables us to test whether either the SCP or contestability/efficiency hypothesis can be rejected or not.

2 See *European Commission* (1997) for a detailed evaluation of the impact of the single market programme on the performance and strategic reaction of the banking sector in European countries.

III. The reduced-form revenue model

The empirical PR model allows of observing and testing the banking industries' market structure. It is a reduced form equation in logarithms, derived from log-linear marginal revenue and cost functions. In the empirical analysis, we use the following operationalisation with a stochastic error term (e):

$$(1) \qquad \ln \mathrm{INTR} = \alpha + \beta \ln \mathrm{INTE} + \gamma \ln \mathrm{PE} + \delta \ln \mathrm{CE} + \varepsilon \ln \mathrm{BSF} + \eta \ln \mathrm{OI} + e$$

where:

INTR = ratio of total interest revenue to the total balance sheet
INTE = ratio of annual interest expenses to total funds
PE = ratio of personnel expenses to the total balance sheet
CE = ratio of physical capital and other expenses to fixed assets
BSF = proxy for bank specific factors
OI = ratio of other income to the total balance sheet

INTE, PE and CE are the unit prices of the inputs of the banks: funds, labour and capital, or proxies of these prices. The H statistic is equal to the sum of the related elasticities: $H = \beta + \gamma + \delta$. The "ratio of personnel expenses to the number of employees" could be a plausible alternative to the "ratio of personnel expenses to the total balance sheet" (PE) included in our estimations. However, the former proxy is only available for a small subset of our observations. The "ratio of physical capital and other expenses to fixed assets" is also a proxy. In particular, the balance sheet item "fixed assets" appears to be unrealistically low for some banks.[3] "Capital expenses" includes the cost of premises, equipment and information technology.

The dependent variable is "ratio of total *interest* revenue to the total balance sheet", as in Molyneux et al. (1994). The choice for taking only the interest part of the total revenue of banks is consistent with the underlying notion of the PR model that financial intermediation is the core business of most banks. However, Shaffer (1982) and Nathan and Neaves (1989) have chosen total revenue as dependent variable.[4] Actually, in recent years, the share of non-interest revenues to total income has increased. We also include the ratio of other income to the total balance sheet (OI) as explanatory variable to account for the influence of

3 However, the exclusion of outliers did not change the estimations remarkably.

4 Further discussion and some empirical evidence regarding the specification of the dependent variable is provided at the end of Section IV.

the generation of other income on the model's underlying marginal revenue and cost functions. Actually, our benchmark, model (2), is a generalisation and nests the models of Molyneux et al. ($\eta = 0$) and Nathan and Neaves ($\eta = -1$):

$$\text{(2)} \qquad \ln \mathrm{INTR} + (-\eta) \ln \mathrm{OI} = \alpha + \beta \ln \mathrm{INTE} + \gamma \ln \mathrm{PE} + \delta \ln \mathrm{CE} + \varepsilon \ln \mathrm{BSF} + e$$

Bank specific factors (denoted by BSF) are other explanatory variables that, at least theoretically, reflect differences in risks, costs, size and structures of banks, descending from the marginal revenue and cost functions underlying equation (1). The risk component can be proxied by the ratio of risk capital or equity to total assets (EQ) and the ratio of loans to total assets (LO). To capture differences in the deposit mix, the ratio of interbank deposits to total deposits (BDEP) or the ratio of interbank demand deposits to total demand deposits (BDD) can be used. Variations in the loan portfolio are mirrored in the ratio of customer loans to total loans (CL). Finally, divergent correspondent activities are taken into consideration when the ratio of cash and due from depository institutions (or banks) to total deposits (CDFB) is included. Total assets (TA) are used as the scaling factor.

We expect a negative coefficient for EQ, because less equity implies more leverage and hence more interest income (Molyneux et al., 1994). A positive parameter for LO is expected, as more loans reflect more potential interest rate income. The coefficient for OI is probably negative as the generation of other income may be at the expense of interest income. Moreover a negative value is in between the extreme cases mentioned above. There are no strong *a priori* expectations regarding the signs of the coefficients of the other explanatory variables. Estimations with these bank-specific variables only yielded satisfactory outcomes in a few instances. Some variables did not have the theoretically expected signs, and were consequently deleted from the specifications. We have also omitted additional variables if their inclusion substantially reduced the number of available observations[5].

To verify whether the competitive structure has changed due to the process of liberalisation and deregulation, we apply model (1) to a pooled cross-section – time-series analysis over the time span 1989 - 96. Unlike Molyneux et al. (1994), we assume that the market structure shifts *gradually* over time. Without this assumption of gradual change,

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[5] A substantially reduced set of observations may yield less representative results.

the results may be implausibly erratic, as found by Molyneux et al. (1994). On the other hand, ignoring these gradual market dynamics may lead to imprecise parameter estimates and biased H statistics, which could in turn result in wrong inferences about the competitive nature of the banking industry. Therefore, we have incorporated a logistic time-curve model into model (1). This way, we explicitly account for possible time variations in the elasticities of the H statistic.

$$\begin{aligned}(3)\qquad \ln \mathrm{INTR} = \alpha + f(\beta,\ \tau, \mathrm{TIME})\,(\ln \mathrm{INTE} + \gamma' \ln \mathrm{PE} + \delta' \ln \mathrm{CE})\\ + \varepsilon \ln \mathrm{BSF} + \eta \ln \mathrm{OI} + e\end{aligned}$$

with $f(\beta, \tau, \mathrm{TIME}) = 1/(1 + (1/\beta - 1)\exp(\tau\,\mathrm{TIME}))$, $\beta \neq 0$,[6] $\gamma' = \gamma / f(\beta, \tau, \mathrm{TIME})$ and $\delta' = \delta / f(\beta, \tau, \mathrm{TIME})$. In this model H is equal to $f(\beta, \tau, \mathrm{TIME})\,(1 + \gamma' + \delta')$. Note that (1) and (3) are identical for $\tau = 0$ and that H increases over time if τ is negative.

IV. Empirical results of the reduced-form revenue model

1. *Estimations for the EU banking industry*

In this section, the original and modified H statistic approach is applied to the EU as a whole. This way, we can investigate the structure of the EU banking industry and the effects of underlying dynamics such as deregulation and liberalisation over time. For the moment, we assume that one EU-wide banking market exists. Data have been collected for the fifteen EU countries.[7] The main source is the International Bank Credit Analysis Ltd (Fitch-IBCA), a London-based bank credit rating agency. Table 1 reports the number of banks and available observations for each country.[8] In principle, individual bank data cover the period 1989 - 96. For the Netherlands, however, the sample period starts in 1992, because no figures were published on interest received and interest paid before 1992. The total number of banks included in the sample is 892, yielding 5259 observation points over the time period considered.

6 If $\beta = 0$ *or* $\beta \approx 0$, the model can be reparameterised, e.g.: $1/(1 + (1/\gamma - 1)\exp(\tau\,\mathrm{TIME}))\,(\ln \mathrm{PE} + \beta' \ln \mathrm{INTE} + \delta' \ln \mathrm{CE})$. Also, if $\beta = 1$, the model needs to be reformulated.

7 Austria, Finland and Sweden are EU Member States since 1 January 1995, i.e. not over the whole time span considered.

8 Note that ignoring observations of non-financial institutions, which also provide financial intermediation on some subdivision of the banking market, does not distort the current analysis, as the actual (overall) competitive conditions are observed directly, irrespective of the providers of intermediation services.

Table 1

Number of banks and observations per country

	Number of banks	Number of observations
Austria	58	399
Belgium	55	316
Denmark	37	240
Finland	10	61
France	89	584
Germany	88	624
Greece	22	131
Ireland	50	64
Italy	92	633
Luxembourg	104	570
Netherlands	45	209
Portugal	38	238
Spain	80	603
Sweden	27	140
United Kingdom	97	447
Total	892	5259

Table 2 shows the pooled regression results of the model applied to observations over the sample period 1989 - 96 of 892 banks from the fifteen EU countries. To test for the robustness of the estimations results, this Table records the outcomes of various model specifications. We will first discuss the preferred specification recorded in Column 1.

The three unit prices of the banks' inputs all contribute significantly to the explanation of the interest revenue of banks. Given the fact that funding is the main factor in the production function of banks, it is hardly surprising that its elasticity, β, is the largest one, followed by the coefficient of labour. The three factor price elasticities decrease only slightly over time,[9] although statistically significant. In the first year (1989), the values of β, γ and δ are 0.675, 0.101 and 0.011, respectively (the logistic curve time-trend effect has been incorporated), whereas these parameters amount to 0.653, 0.098 and 0.011, respectively, in the last year (1996). Given the recent deregulation and liberalisation in the

9 The change in the time trend coefficient of the logistic curve, τ, is significant.

EU, it is surprising that H declines (from 0.787 to 0.762), rather than increases. Apparently, opposite influences shaping the competitive environment have been at work. The effect of deregulation and liberalisation, on the one hand, and the effect of an increase in the degree of concentration, on the other, have probably partly offset each other.[10] Formal tests reveal that H differs significantly from both 0 and 1,[11] providing evidence that a certain degree of monopolistic competition in the European banking market is present. Hence, the hypothesis of either monopoly or perfect competition must be rejected. This finding agrees with our earlier expectation that competitive interactions between banks can best be characterised by monopolistic competition.

The degree of competition is comparable to that found by Nathan and Neave (1989) for the Canadian bank market and is higher than found by Shaffer and Lloyd-Williams et al. (1991) for, respectively, the New York and Japanese bank markets. H is also much higher than found by Molyneux et al. (1994) for *the averages* over 1986 - 1989 of bank markets in the major European countries.[12] As mentioned earlier, their yearly figures deviated substantially from these averages. The level of H compares to that for Finland found by Vesala (1995).

Apart from common factors, the structure of the banking markets in individual EU countries also – still – depends on numerous country-specific features, e.g. national institutions, the degree of government intervention, the sophistication of the financial system, etc. Moreover, national balance sheet figures are not always perfectly comparable due to existing differences in national accounting practices and definitions. To take these country-specific characteristics into account, we have added dummy variables for the various countries.

The country-specific dummies also contribute significantly to the explanation of the dependent variable. More importantly, the value of the dummies varies considerably across countries and ranges between 0.49 for France and 1.44 for Denmark. A formal test indeed strongly rejects the hypothesis that the dummy parameters are all equal to each other.[13] This points to structural differences between national banking

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10 Changes over time in the yield curve may also have affected the time trend shift. This could be avoided by including a measure of the spread as additional explanatory variable.

11 The χ^2-statistic of the Wald test of $1/(1 + (1/\beta - 1)\exp(\tau 4.5)(1 + \gamma' + \delta') = 1$ is 533.6 (probability value: 0.0000 %).

12 *Molyneux* et al. (1994) found averages for France, Germany, Italy, Spain and the UK of, respectively, 0.61, 0.31, -0.63, 0.13 and 0.07. The overall average is 0.18.

Table 2

Empirical results of the reduced-form revenue model for the EU

		Model (3)						Model (1)	
				Extra variables		Weighted Coeff.			
Column		1		2		3		4	
INTE	β	0.68	*92.1*	0.65	*94.0*	0.83	*180.0*	0.67	*93.0*
PE	γ	0.15	*32.2*	0.13	*27.1*	0.09	*42.1*	0.10	*32.5*
CE	σ	0.02	*2.5*	0.03	*5.0*	-0.02	*4.2*	0.01	*2.4*
TIME	τ[a]	1.38	*4.6*	1.09	*4.1*	2.00	*9.1*		
LO				0.09	*26.9*				
TA				-0.02	*10.0*				
Austria		1.23	*45.2*	0.86	*28.2*	0.91	*48.0*	1.21	*44.9*
Belgium		0.59	*20.0*	0.27	*9.1*	0.34	*20.8*	0.57	*19.7*
Denmark		1.44	*52.4*	1.04	*34.2*	0.94	*52.2*	1.43	*52.2*
Finland		1.23	*32.5*	0.81	*20.9*	0.87	*40.0*	1.21	*32.1*
France		0.49	*18.0*	0.17	*5.9*	0.34	*21.0*	0.47	*17.7*
Germany		0.58	*22.9*	0.31	*10.9*	0.41	*25.1*	0.57	*22.7*
Greece		0.58	*17.5*	0.31	*9.3*	0.33	*16.3*	0.56	*17.0*
Ireland		0.66	*18.6*	0.35	*9.6*	0.61	*26.2*	0.64	*18.2*
Italy		1.28	*48.4*	0.90	*30.4*	0.86	*50.0*	1.26	*48.2*
Luxembourg		0.54	*19.4*	0.33	*11.7*	0.40	*23.7*	0.52	*19.0*
Netherlands		0.73	*25.6*	0.40	*13.2*	0.59	*37.2*	0.71	*26.2*
Portugal		0.63	*21.5*	0.33	*11.3*	0.48	*27.2*	0.61	*21.1*
Spain		0.72	*28.0*	0.43	*16.0*	0.46	*28.1*	0.70	*27.9*
Sweden		1.42	*41.5*	1.01	*27.3*	0.92	*43.3*	1.39	*41.1*
UK		0.52	*18.1*	0.27	*9.1*	0.43	*25.6*	0.50	*17.7*
Adj. R-squared		0.711		0.750		0.998		0.710	
H (1989/1996)		0.79-0.76		0.75-0.73		0.88-0.86		0.78	
# of observations		5224		5207		5224		5259	

Explanatory note: t-values are presented behind the coefficients in italics.
[a] All coefficients τ are multiplied by 100 for the sake of presentation.

sectors and corroborates the findings of Vives (1991). Interestingly, the dummies are relatively high for countries that have experienced serious problems in their banking systems. Around 1990, the Nordic countries witnessed severe banking crises due to the collapse of the property markets, where they had accumulated large exposures after the deregulation of traditional banking markets. For Italy, this is less clear but it may relate to problems with *Banca di Napoli*.

Column 4 of Table 2 shows the estimation results for model (1), i.e. without the logistic time curve, to verify how the inclusion of the curve has influenced the parameter estimates. It can be concluded that the results hardly change. The H statistic value is 0.78 (constant over time) according to model (1) and decreases slightly from 0.79 to 0.76 according to model (3).

To test whether bank-specific factors are unduly omitted, we experimented with five additional explanatory variables in model (3). The ratio of loans to total assets (LO), reflecting risk, has a significantly positive coefficient, as expected. Cash and due from banks to customer and short-term funding (CDFB), reflecting correspondent activity, is not significant, whereas the ratio of equity to total assets (EQ), also representing risk, and the ratio of other income to the total balance sheet (OI) have the wrong sign. The variable total assets (TA) takes scale effects into account. Other potential explanatory variables caused estimation problems and are dropped henceforth.[14] We also omitted insignificant variables and variables with wrong signs. Therefore, only the expression with the loan ratio and total assets variable is recorded in Column 2 in Table 2. In general, the additional explanatory variables do not substantially affect the components β, γ and δ of the H statistic.

We also ran regressions where banks are weighted according to their size in terms of total assets.[15] This approach is justified by the obvious fact that small banks are less important in the intermediation process than the large ones.[16] Consequently, the competitive behaviour of large

13 A reduction from 15 country dummies to one dummy for Scandinavian countries, one for Italy and Austria and one for the remaining countries is also rejected.

14 For instance, data for an explanatory variable are not available for at least one country.

15 Dependent and explanatory variables (including dummies) of model (3) are all multiplied by "total assets". Total assets for each country are normalised, to correct for the currency unit. The coefficients of the country-specific dummy variables are corrected for this currency unit related normalisation.

16 Weighting can also be justified on econometric grounds. Based on the central-limit theory, large banks behave more "normally" than smaller ones. We found

banks should have a greater impact on the estimation outcomes in order to obtain a more representative and realistic picture of the actual competitive conditions prevailing in the banking market. Column 3 of Table 2 shows the results for model (3) with weighted data. The elasticity of funding costs (β) is much higher in the weight-based model (0.83) than in the standard model (0.68). Its effect on H is slightly offset by lower values of the other cost components. On balance, H rises in the weight-based model to 0.88 in 1989 and 0.86 in 1996 (from, respectively, 0.79 and 0.76 in the standard model). From this shift in H it can be concluded that the market environment of the larger banks is more competitive. Presumably, larger banks are relatively more active in international wholesale markets, where competition is generally assumed to be more fierce (Gilibert and Steinherr, 1989). In banking, relatively smaller competitive pressures exist in the retail market and for small-to-medium sized firms in the corporate banking market. The coefficients of the country-specific dummy variables in the weighted model are comparable to those of the standard model.[17]

Before concluding, two issues remain to be investigated. Firstly, as elaborated in the literature, a critical feature of the H statistic is that the PR-approach must be undertaken on the basis of observations that are in long-run equilibrium. An equilibrium test uses the fact that in competitive capital markets, risk-adjusted rates of return will be equalised across banks. In such a case, the rates of return will not be correlated with input prices. An equilibrium test is provided by model (1), after replacement of the dependent variable by rate of return on total assets (ROA) or equity (ROE). $H = 0$ would then indicate equilibrium, whereas $H < 0$ would point to disequilibrium. Using ROE, we find that the hypothesis of equilibrium ($H = 0$) cannot be rejected at the 95% significance level.[18] This justifies the applied methodology.

Secondly, the PR methodology refers to the elasticities of the reduced form or equilibrium revenues with respect to changes in input prices. Like Molyneux et al. (1994), we use revenues scaled by the total balance sheet value as dependent variable in the estimation to generate estimates of the H statistic. Vesala (1995) points out that scaling transforms the rev-

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empirical evidence that disturbance terms of big banks are indeed (somewhat) smaller than those of smaller banks. Weighted estimation is then more efficient.

17 All the dummy coefficients are lower, due to a decline in the implicit overall intercept.

18 The χ^2-statistic of the Wald test of $1/(1 + (1/\beta - 1)\exp(\tau 4.5)(1 + \gamma' + \delta') = 0$ is 1.75 (probability value: 18.6%).

enue equation into a pricing equation. The comparative static properties of the reduced-form revenue equation that guide our judgement of competitive conditions do not necessarily hold as such for pricing equations. Vesala (1995) suggests re-estimation using revenues as the dependent variable to see whether the conclusions remain the same. Following this suggestion and in line with Shaffer (1982) and Nathan and Neave (1989), we have (re-)estimated our model with the (unscaled) total revenue as dependent variable. For the total-revenue model, the value for H is decreasing from 0.68 (in 1989) to 0.66 (in 1996), instead of decreasing from 0.79 to 0.76 as in the scaled-interest-income model. Hence, this change in the dependent variable does not affect the conclusions: in any case H differs significantly from both 0 and 1. Actually, Nathan and Neave (1989) take total revenue *corrected for loan losses* as dependent variable. Also, the impact of this correction appears to be virtually negligible.

The overall conclusion is that different specifications barely affect the parameter values underlying the H statistic. For the EU as a whole, the banking market is characterised by monopolistic competition. In spite of the deregulation and liberalisation of the EU banking markets over 1989 - 96, we find hardly any evidence of increasing competition over the years. The divergent country-specific dummies demonstrate that the hypothesis of a single European banking industry must be strongly rejected.

2. *Empirical results for the individual countries*

Judging by the divergent values of the country-specific dummies emerging from the EU-wide analysis, national banking systems clearly exhibit structural differences. This calls for an empirical assessment of the competitive environment encountered by national banking sectors. A closer look at the competitive behaviour of the national banking industry also enables us to explore the relationship between concentration and competition. For these reasons, we also apply the above PR analysis to the banking sectors in all individual EU countries separately.

Tables A.1 - A.15 in the Appendix present the empirical results for each country. The first two columns of each Table show the preferred version of the *standard* country model. Where the time trend was insignificant, the logistic time-trend model was deleted. Where sufficient[19]

[19] Often, additional variables were available for a limited set of banks only, so that inclusion of these variables would have strongly reduced the number of observations. Therefore, those variables were not included.

data were available, additional explanatory variables were initially included to check for differences in risk and other bank-specific characteristics. However, these additional variables were dropped from the equation if their parameters were wrongly signed or insignificant.[20] The loan ratio (LO) and the scale variable total assets (TA) proved to be the most successful as explanatory variables. In half the country models, they were significant. The other explanatory variables are used less frequently. For the sake of comparison, the "standard" model without additional explanatory variables, such as in the EU section, is shown in the two centre columns of each Table. Differences in the results may be due to the various specifications of the model, but also to different sets of available observations.

Table 3

***H* statistics for the EU countries (1989 - 96)**

	Standard model	Weighted model
Austria	0.80 - 0.77	0.87
Belgium	0.92^a	$0.94 - 0.98^a$
Denmark	0.10 - 0.48	0.36 - 0.61/0.83
Finland	0.68 - 0.56	0.81 - 0.72
France	0.91	$0.99 - 0.96^a$
Germany	0.84 - 0.86	0.92 - 0.91
Greece	0.92^a	0.94^a
Ireland	0.39	0.75 - 0.69
Italy	0.88 - 0.94	0.96^b
Luxembourg	0.92	0.86
The Netherlands	0.66	0.73 - 0.70
Portugal	0.79 - 0.73	0.86
Spain	0.60 - 0.55	0.83 - 0.78
Sweden	0.59	0.78 - 0.74
UK	0.72	0.76 - 0.70

Explanatory notes: ^{a} $H = 1$ not rejected at the 95% level of confidence. ^{b} $H = 1$ not rejected at the 99% level of confidence.

20 Insignificant explanatory variables were dropped, as extra variables often imply some loss of observations.

Table 3 presents the values of H for all countries. According to the standard model (left column), H appears to be high, between two-third and one, in most countries. Only for two smaller countries, lower values for H are observed: rising from 0.10 in 1989 to 0.48 in 1996 in Denmark and 0.39 in Ireland. The hypothesis $H = 0$, implying monopoly, a perfectly colluding oligopoly or a conjectural variations short-run oligopoly, is rejected for each country. On the other hand, the hypothesis of perfect competition ($H = 1$) cannot be rejected at the 95% confidence level for two countries only, *viz.* Belgium and Greece. Hence, the banking markets in separate EU countries can generally be classified as monopolistic competition, much closer to perfect competition than to monopoly. The competitive environment has also hardly changed over the years considered. For three countries, H rises somewhat over time and for four others H falls slightly. An exception to this general picture is Denmark, where a sharp increase in H has taken place. In a sense, these findings conflict with the widely accepted notion that the EEC's 1992 Second Banking Directive brought about substantial cross-country financial integration and helped to change the competitive environment in European banking. According to our estimates, competition was also already rather fierce in the late 1980s and has hardly increased since then.

The first three columns of Table 4 present the constituent parts of H of the standard country model. The coefficients of interest rate expenditures (β), with values between 0.6 and 0.8, contribute by far the most to H in all countries, with the exception of Denmark (0.06[21]). The coefficients of personnel expenses (γ) also contribute significantly and substantially to H with values ranging mainly between 0.10 and 0.15. Outliers are Denmark (0.39) and Italy (0.32), on the one hand, and France, Ireland and the Netherlands (0.02 to 0.03), on the other. The coefficients of physical capital expenses (δ) vary in sign and magnitude. On average, their contribution to H is negligible.

As in the EU-wide analysis, we also carried out the same exercise with larger weights for larger banks. The two last columns of Tables A.1 - A.15 give the results of these weighted estimations. The right column of Table 3 in the main text displays the corresponding H statistics. Consistent with our earlier findings, the H statistics are higher than those resulting from the estimation of the standard model for all countries, except Luxembourg. In particular for Denmark and Ireland,

[21] The coefficient of interest rate expenditures, including the logistic time trend effect, ranges from 0.08 in 1989 to 0.38 in 1996.

Table 4

Elasticities of the unit prices

	Standard model				Weighted model			
	β	γ	δ	$\beta+\gamma+\delta$	β	γ	δ	$\beta+\gamma+\delta$
Austria	0.72	0.13	-0.01	0.84	0.76	0.11	0.00	0.87
Belgium	0.74	0.08	0.10	0.92	0.91	0.12	-0.09	0.94
Denmark	0.06	0.39	-0.12	0.33	0.72	0.15	-0.04	0.83
Finland	0.62	0.09	0.04	0.75	0.65	0.18	0.10	0.93
France	0.84	0.03	0.04	0.91	0.92	0.03	0.05	1.00
Germany	0.72	0.16	0.00	0.88	0.83	0.10	0.01	0.94
Greece	0.70	0.14	0.08	0.92	0.73	0.20	0.01	0.94
Ireland	0.65	0.03	-0.29	0.39	0.58	0.16	0.13	0.87
Italy	0.63	0.32	0.06	1.01	0.77	0.16	0.03	0.96
Luxembourg	0.82	0.08	0.02	0.92	0.81	0.05	0.00	0.86
The Netherlands	0.65	0.02	-0.01	0.66	0.69	0.08	-0.02	0.75
Portugal	0.77	0.12	-0.07	0.82	0.80	0.06	0.00	0.86
Spain	0.49	0.14	0.11	0.74	0.69	0.21	0.01	0.91
Sweden	0.47	0.08	0.04	0.59	0.80	0.02	0.04	0.86
UK	0.63	0.04	0.05	0.72	0.65	0.07	0.10	0.82
Average	0.63	0.12	0.00	0.76	0.75	0.11	0.02	0.89
Standard deviation	*0.20*	*0.11*	*0.10*	*0.21*	*0.09*	*0.06*	*0.05*	*0.06*

the shift which reflects the difference in behaviour between smaller and larger banks, is remarkable. H is below 1 for all countries. The hypothesis $H = 0$ is rejected for all countries, which excludes monopoly, whereas perfect competition ($H = 1$) cannot be rejected at the 95 % level of confidence for Belgium, France, Greece, and Italy. The banking markets in the latter four countries are characterised by perfect competition ($H = 1$) or monopolistic competition ($H \leq 1$). For the other countries, the banking market is characterised by monopolistic competition. The upward shift in H is mainly due to a higher value of the parameter β, averaging 0.75 in the weighted model versus 0.63 in the standard model (Table 4). Compared to the original model, the values of the price elasticities β, γ and δ also show a remarkable convergence. This is confirmed by the relatively low standard deviations of these coefficients across countries in the weighted version of the model (last row of Table 4). The fall in the standard deviation is most prominent for the sum of the elasticities:[22] from 0.21 to 0.06. This remarkable improvement in the stability of the elasticities suggests that the weighted model produces more reliable results.

As discussed before, an economic interpretation of this upward shift in H could be that larger banks operate in relatively more competitive markets (e.g. wholesale markets) than smaller banks. At first sight, this conflicts with the widespread belief that banks with large market shares may exert stronger market power, resulting in a more monopolistic market structure. Plausible explanations are that smaller banks tend to have a stronger position in local and specialised markets. Their core business could be to attract deposits from economic agents in the region where they are located. Small banks are also expected to have more knowledge about local firms. Consequently, smaller banks are also, or in some cases perhaps even slightly better, able to acquire some market power than their internationally active large counterparts.[23] At this point, one should also remember that the PR model concentrates on the competitive conditions faced by the entire banking industry, irrespective of whether national banks are more domestically or more internationally oriented. Hence, the distinction between national and cross-border banking activities plays a subordinate role in the PR framework. Finally, with regard to the effect of the time trend on H, a significant (but minor) rise in H is observed for two countries only, whereas a slight (but statistically significant) decrease is found for five countries.

So far, we have hardly paid attention to the underlying causes of the different H statistics across countries. Although the statistics point to the existence of monopolistic competition, in reality one can discern differences in the degree of competition. In a sense, the results may even seem somewhat counterintuitive. Indeed, banking sectors that are generally considered to suffer from a lack of competition and excess capacity, namely the banking systems of the southern European countries and Germany, behave rather competitively. Here, it should be noted that some economists mix up the concepts of efficiency and competition in banking or use them interchangeably, which definitely obscures the pic-

[22] Note that $\beta+\gamma+\delta$ deviates from H, if the model contains a time trend.

[23] As predicted by the theory on relationship banking (*Conigliani* et al., 1997).

ture and causes a lot of confusion. Inefficient banking systems can be profitable and competitive. On the other hand, a high degree of competitiveness gives no guarantee that bank sectors work efficiently (Molyneux *et al.*, 1996). Explaining the underlying causes of the different H statistics across countries would require a detailed analysis of the institutional, economic, regulatory and political features of national banking systems, which is beyond the scope of this paper. Nevertheless, the next section seeks to lift a corner of the veil.

V. Competition and concentration

This section investigates the relationship between the degree of competition, i.e. the value of H and the level of concentration in banking systems. As discussed in Section 2 the nature of this connection is not straightforward *a priori*. The assets of the five biggest banks as a percentage of total assets (C5) are used as a measure of concentration. The World Bank provides 1995 figures of C5, based on IBCA (Demirgüç-Kunt and Huizinga, 1998); for earlier years the IBCA database is less complete. Table 5 presents the World Bank concentration index C5 and the H statistics from the standard version (HS) of our model. Where a time trend is included in the model, the 1995 values of H have been taken.

Chart 1 shows a scatter of C5 and *HS* for each country, as well as an estimated linear regression line through these data points. The downward sloping regression line implies that higher concentration is indeed accompanied by a lower degree of competition, which supports the structure-conduct-performance hypothesis. The negative slope of the regression line has a t-value of 2.1, implying that the inverse relationship between concentration and competition is significant at the 95% level of confidence. A weighted regression line, with the number of banks from Table 1 as weights (assuming that larger countries with more banks are more important), yields a more significant negative slope with a t-value of 2.6. Despite some well-known conceptual difficulties, there appears to be evidence for the often-assumed negative effect of concentration on competition. This relationship also partly explains the observed gradations in monopolistic competition (i.e. variations in H) across countries.

Additional information about the relationship between competition and concentration is provided by rank correlation tests, which have the advantage that they are not affected by specification choices. The correlation coefficient between the rank numbers of two series (here HS and C5) is known as Spearman's rank correlation coefficient, which is sym-

Table 5

Concentration ratio and competition in EU countries (1995)

	C5	HS
Austria	0.75	0.77
Belgium	0.46	0.92
Denmark	0.77	0.43
Finland	0.70	0.58
France	0.48	0.91
Germany	0.50	0.86
Greece	0.70	0.92
Ireland	0.75	0.39
Italy	0.27	0.93
Luxembourg	0.30	0.92
The Netherlands	0.84	0.66
Portugal	0.32	0.74
Spain	0.50	0.56
Sweden	0.41	0.59
UK	0.39	0.72

metrically distributed with zero mean and variance $1/(m-1)$, where m denotes the number of observations (Theil, 1971). For HS this coefficient is -0.491, whereas the standard deviation is 0.267. This also points to a negative effect of concentration on competition.

Kendall (1962) suggests an alternative procedure, which consists of counting the number of inversions in the second ranking (e.g. H) compared to the first one (e.g. C5).[24] The number of inversions, v, ranges from 0 to $m(m-1)/2$. Thus, $1-4v/(m(m-1))$ is a rank correlation coefficient, which ranges from –1 to 1. Its distribution is symmetric around zero and its variance is $2(2m+5)/(9m(m-1))$. For $m > 10$ a normal approximation can be used (Kendall and Stuart, 1967). For *HS* the value of v is 31, and, so, the rank correlation coefficient is 0.410, whereas the standard deviation is 0.192. Again, a significant negative relationship between concentration on competition is found.

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[24] If the ranking of C5 is (after re-arrangement) 1, 2, 3, and the corresponding ranking of H is 3, 2, 1, the number of inversions is three: 3 is before 2 and 1, and 2 is before 1.

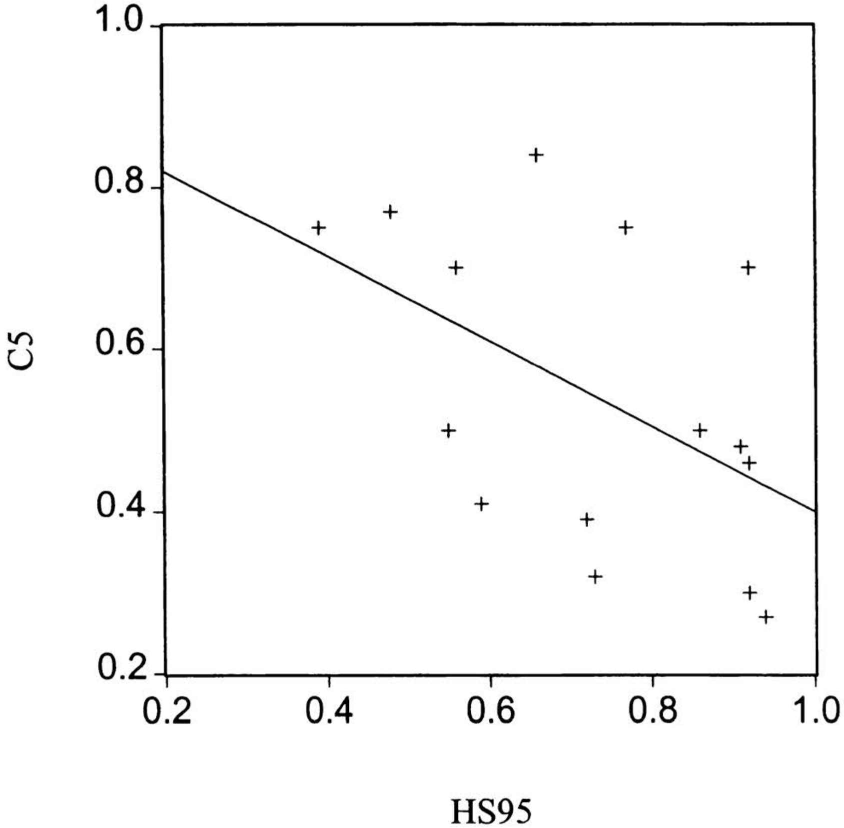

Figure 1: Relationship between concentration (C5) and competition (H) in 1995

In summary, the results for the EU indicate that higher concentration is accompanied by a lower degree of competition, which supports the SCP hypothesis. These tests of the SCP paradigm have raised objections regarding the fact that the geographical unit underlying the concentration measure does not coincide with the banking markets and that the influence of non-bank financial institutions is neglected. These objections are only partly valid, though. As mentioned earlier, the concept of competition does not pertain to national markets, but to markets in which national banks operate. The concentration index does not necessarily coincide with domestic markets, too. The overall indices, H and C5, do not allow for a distinction between different segments of the banking markets, such as retail or wholesale markets. Therefore, our results are broad in nature, too.

The *H* statistic measures the degree of competition directly and does not require observations of non-bank financial institutions as well. The concentration index indeed neglects non-banks, but this shortcoming mainly refers some market segments such as mortgage lending and does not substantially affect the overall conclusions. Probably, the SCP hypothesis holds more for local than for international markets. Even if concentration in local markets adversely affects competition, supervisory authorities may well tolerate a certain degree of it, if this is required to obtain the critical mass which is necessary to serve international clients or to survive on relatively competitive banking markets.

VI. Concluding remarks

In this paper, we have developed an amended version of the traditional Panzar-Rosse methodology to assess the degree of competitiveness in the banking industry of the EU as a whole and of individual countries. One of the innovations concerns the inclusion of a logistic time-curve model in the original model to capture the potential effects of deregulation and liberalisation, concentration and shifts in the yield curve adequately. Contrary to most existing studies, we have used a pooled cross-section – time-series analysis over the time span 1989 - 1996. Moreover, we have also run regressions where greater weights are attached to larger banks. This approach does justice to the fact that large banks play a greater role in the financial intermediation process.

Compared to Molyneux et al. (1994), we obtain more plausible and fairly stable results. The competitive nature of banking markets appears to change only very gradually over time and can be qualified as monopolistically competitive in most countries. Some confidence can be placed in – the robustness of – these results, because different specifications point in the same direction. Our measure of competitiveness displays hardly any dynamics. This challenges the conventional view that the process of deregulation and liberalisation of financial services triggered by the EEC's 1992 Second Banking Co-ordination Directive would increase competition in the EU banking, among other things. At the time of the adoption of this Directive and the accompanied formulation of additional policy measures, competition was already rather fierce. Moreover, our estimations also reject the hypothesis that national banking sectors in the EU are identical. Hence, European countries enter EMU with different national banking systems.

In a similar vein, it seems now widely accepted that the establishment of EMU will reinforce already prevailing trends in European banking. Undoubtedly, EMU will bring about major changes for the European banking industry. However, the general notion that EMU will both increase the degree of competition and concentration is not supported by our empirical study. On the contrary, if EMU leads to further rationalisation and consolidation in the banking industry, one cannot exclude that the overall environment will become less competitive. This assessment is based on indications of a negative correlation between the degree of competition and concentration, which is consistent with a weak version of the traditional SCP-paradigm. This is an important message for European policy-makers and banking supervisors if they want to maintain the present high level of competition. If so, any increase in concentration, e.g. through mergers or acquisitions, must be carefully evaluated.

Appendix

Table A.1

Empirical results for Austria

	Standard				Weighted	
	Coeff.	t-values	Coeff.	t-values	Coeff.	t-values
INTE	0.72	*55.9*	0.70	*54.0*	0.76	*158.0*
PE	0.13	*10.2*	0.11	*13.9*	0.11	*33.4*
CE	-0.01	*1.1*	-0.03	*4.9*	-0.00	*0.1*
TIME[a]	1.67	*1.9*				
LO	0.03	*3.5*				
BDEP	-0.02	*4.7*			-0.03	*6.6*
BDD					-0.03	*4.4*
CDFB					0.00	*3.2*
TA	-0.01	*2.8*				
Intercept	1.22	*20.0*	1.39	31.3	1.37	*39.3*
Adj. R-squared	0.919		0.882		1.000	
Number of observations	332		399		322	
H (1989/1996)	0.80	0.77	0.78		0.87	

Explanatory notes: $H = 0$ and $H = 1$ are rejected for all equations (level of confidence 99.9%). [a] All coefficients τ in this and later Tables are multiplied by 100 for the sake of presentation.

Table A.2

Empirical results for Belgium

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.74	*25.3*	0.76	*20.7*	0.91	*89.3*
PE	0.08	*9.3*	0.14	*9.4*	0.12	*5.9*
CE	0.10	*5.7*	0.10	*3.7*	-0.09	*5.8*
TIME			29.5	*2.6*	-9.31	*2.1*
OI					-0.09	*5.8*
LO	0.09	*10.4*				
EQ	-0.10	*7.5*				
TA	-0.03	*5.3*				
Intercept	-0.20	*1.8*	0.15	*1.1*	0.70	*13.1*
Adj. R-squared	0.751		0.619		0.999	
Number of observations	314		316		316	
H (1989/1996)	0.92		0.93	0.89	0.94	0.98
$H = 1$[a]	3.55[b]	(5.9)	3.15[b]	(7.6)	0.76[b]	(38.3)

Explanatory notes: $H = 0$ is rejected for all equations at the 99% level of confidence. [a] For $H = 1$ the values of the F-statistic are shown and, between brackets, the probability of a wrongly rejected null hypothesis, in % (2.8% means null hypothesis rejected at the 5% level of significance and not rejected at the 1% level of significance). [b] $H = 1$ is not rejected at the 95% level of confidence. [c] $H = 1$ rejected at the 5% level of significance but not rejected at the 1% level of significance.

Table A.3
Empirical results for Denmark

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.06	*3.1*	0.03	*2.6*	0.72	*52.6*
PE	0.39	*10.4*	0.60	*6.2*	0.15	*15.5*
CE	-0.12	*3.3*	-0.05	*0.8*	-0.04	*3.8*
TIME	-28.58	*5.0*	-30.67	*4.6*		
CDFB	-0.02	*2.5*			-0.01	*3.7*
LO	0.22	*12.6*			0.32	*8.0*
TA	-0.04	*7.9*				
Intercept	1.55	*21.5*	2.40	*61.4*	0.32	*2.8*
Adj. R-squared	0.765		0.553		1.000	
Number of observations	236		240		236	
H (1989/1996)	0.10	0.48	0.06	0.42	0.83	

Explanatory note: $H = 0$ and $H = 1$ rejected for all equations (level of confidence 99.9%).

Table A.4
Empirical results for Finland

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.62	*12.3*	0.62	*12.0*	0.65	*29.5*
PE	0.09	*2.6*	0.07	*2.2*	0.18	*8.5*
CE	0.04	*0.7*	0.05	*0.9*	0.10	*5.4*
TIME	6.58	*2.8*	5.39	*2.5*	4.46	*2.7*
LO	0.09	*1.7*			0.16	*3.4*
EQ	-0.07	*2.2*				
TA					-0.00	*6.7*
Intercept	0.90	*3.8*	1.04	*7.5*	0.66	*3.4*
Adj. R-squared	0.793		0.777		0.999	
Number of observations	61		61		61	
H (1989/1996)	0.68	0.56	0.68	0.57	0.81	0.72

Explanatory note: $H = 0$ and $H = 1$ are rejected for all equations (level of confidence 99.9%).

Table A.5

Empirical results for France

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.84	*36.9*	0.82	*34.0*	0.92	*73.4*
PE	0.03	*2.3*	0.11	*11.0*	0.03	*4.4*
CE	0.04	*2.6*	0.02	*1.4*	0.05	*3.8*
TIME					4.12	*3.9*
LO	0.09	*9.6*			0.04	*5.8*
CDFB	-0.03	*3.2*			-0.20	*18.0*
CL					-0.02	*4.5*
BDD					-0.04	*5.7*
TA	-0.10	*6.7*				
Intercept	0.14	*1.3*	0.13	*1.4*	0.76	*9.7*
Adj. R-squared	0.743		0.676		0.998	
Number of observations	570		584		521	
H (1989/1996)	0.91		0.95		0.99	0.96
$H = 1$[a]	7.4[c]	(0.6)	0.3[b]	(60.0)	0.1[b]	(79.1)

Explanatory notes: $H = 0$ is rejected for all equations (level of confidence 99.9%). [a] For $H = 1$ the values of the F-statistic are shown and, between brackets, the probability of a wrongly rejected null hypothesis, in %. [b] $H = 1$ is not rejected at the 95% level of confidence. [c] $H = 1$ is rejected at the 99% level of confidence.

Table A.6
Empirical results for Germany

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.72	*37.4*	0.69	*29.8*	0.83	*93.9*
PE	0.16	*11.5*	0.17	*18.6*	0.10	*15.1*
CE	0.00	*0.0*	-0.05	*3.8*	0.01	*2.0*
TIME	-1.49	*2.2*	-2.52	*2.8*	0.64	*2.2*
OI	-0.02	*5.4*			-0.02	*11.0*
LO	0.11	*14.4*				
EQ	-0.05	*2.9*				
BDD	-0.18	*13.0*			-0.18	*25.5*
Intercept	0.86	*10.7*	0.67	*9.7*	1.17	*33.5*
Adj. R-squared	0.790		0.622		1.000	
Number of observations	595		624		595	
H (1989/1996)	0.84	0.86	0.78	0.81	0.92	0.91

Explanatory note: $H = 0$ and $H = 1$ rejected for all equations (level of confidence 99.9%).

Table A.7
Empirical results for Greece

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.70	*17.9*			0.73	*24.1*
PE	0.14	*7.4*			0.20	*12.2*
CE	0.08	*4.2*			0.01	*0.4*
TIME						
Intercept	0.12	*0.9*			0.10	*0.9*
Adj. R-squared	0.818				0.999	
Number of observations	131				131	
H (1989/1996)	0.92				0.94	
$H = 1^a$	0.1^b	(76.3)			0.3^b	(85.2)

Explanatory notes: $H = 0$ is rejected for all equations (level of confidence 99.9%). [a] For $H = 1$ the values of the F-statistic are shown and, between brackets, the probability of a wrongly rejected null hypothesis, in %. [b] $H = 1$ is not rejected at the 95% level of confidence.

Table A.8

Empirical results for Ireland

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.65	*15.6*	0.04	*1.0*	0.58	*26.9*
PE	0.03	*1.0*	0.01	*0.2*	0.16	*7.1*
CE	-0.29	*3.8*	-0.48	*15.3*	0.13	*2.7*
TIME			-59.82	*3.5*	2.59	*4.7*
BDD					-0.03	*2.1*
LO					0.12	*2.3*
Intercept	2.20	*5.6*	2.39	*24.5*		
Adj. R-squared			0.907		1.000	
Number of observations			64		59	
H (1989/1996)	0.39		0.03	0.43	0.75	0.69
$H = 0$[a]	17.4[c]	(0.0)	3.1[b]	(7.7)	246.8[c]	(0.0)
$H = 1$[a]	41.4[d]	(0.0)	26.5[d]	(0.0)	41.9[d]	(0.0)

Explanatory notes: [a] For $H = 1$ the values of the F-statistic are shown and, between brackets, the probability of a wrongly rejected null hypothesis, in %. [b] $H = 0$ is not rejected at the 95% level of confidence. [c] $H = 0$ is rejected (level of confidence 99.9%). [d] $H = 1$ is rejected (level of confidence, respectively, 99.9% (twice), 99% and 99.9%).

Table A.9

Empirical results for Italy

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.63	*26.6*	0.68	*29.9*	0.77	*64.7*
PE	0.32	*16.4*	0.20	*18.9*	0.16	*19.6*
CE	0.06	*3.7*	0.08	*5.6*	0.03	*7.7*
TIME	-2.94	*2.7*	-2.18	*4.3*		
LO	0.12	*8.6*			0.05	*3.4*
CDFB	-0.02	*5.4*			-0.02	*7.2*
BDD					-0.02	*2.5*
TA	-0.03	*8.1*				
CL					0.02	*3.3*
Intercept	1.20	*13.6*	1.16	*19.0*	0.94	*9.7*
Adj. R-squared	0.747		0.654		1.000	
Number of observations	601		633		550	
H (1989/1996)	0.88	0.94	0.89	0.95	0.96	
$H = 1^{a}$	10.3^{b}	(0.1)	7.4^{b}	(0.6)	5.5^{c}	(2.0)

Explanatory notes: $H = 0$ rejected for all equations (level of confidence 99.9%). [a] For $H = 1$ the values of the F-statistic are shown and, between brackets, the probability of a wrongly rejected null hypothesis, in %. [b] $H = 1$ is rejected (level of confidence, respectively, 99.9%, 99% and 99.9%). [c] $H = 1$ is rejected at the 5% level of significance, but not rejected at the 1% level of significance.

Table A.10

Empirical results for Luxembourg

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.82	*77.0*	0.80	*67.8*	0.81	*103.5*
PE	0.08	*11.1*	0.06	*7.6*	0.05	*20.1*
CE	0.02	*2.6*	-0.01	*0.8*	0.00	*0.5*
TIME			-2.09	2.4		
CDFB	-0.02	*5.0*				
TA	0.03	*8.1*				
Intercept	0.14	*3.2*	0.38	*8.9*	0.38	*19.8*
Adj. R-squared	0.934		0.894		0.999	
Number of observations	433		570		570	
H (1989/1996)	0.92		0.84	0.86	0.86	

Explanatory note: $H = 0$ and $H = 1$ rejected for all equations (level of confidence 99.9%).

Table A.11

Empirical results for the Netherlands

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.65	*30.0*	0.71	*35.9*	0.69	*27.2*
PE	0.02	*3.4*	0.03	*2.9*	0.08	*13.6*
CE	-0.01	*1.3*	-0.03	*2.2*	-0.02	*1.1*
TIME			4.87	*3.0*	2.72	*4.2*
OI	-0.07	*5.1*				
Intercept	0.62	*8.9*	0.81	*12.4*	0.83	*13.4*
Adj. R-squared	0.879		0.869		0.999	
Number of observations	208		209		209	
H (1992/1996)	0.66		0.70	0.65	0.73	0.70

Explanatory note: $H = 0$ and $H = 1$ are rejected for all equations at the 1% level of significance.

Table A.12

Empirical results for Portugal

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.77	*31.4*	0.75	*31.5*	0.80	*33.9*
PE	0.12	*6.7*	0.13	*7.0*	0.06	*3.5*
CE	-0.07	*4.1*	-0.06	*3.5*	-0.00	*0.1*
TIME	4.29	*5.1*	3.93	*4.8*		
LO	0.06	*2.7*			0.10	*4.5*
EQ	-0.04	*4.0*			-0.04	*3.4*
BDD					-0.08	*6.0*
CDFB					0.02	*3.2*
Intercept	0.61	*6.1*	0.75	*9.2*	0.51	*4.2*
Adj. R-squared	0.850		0.837		0.999	
Number of observations	238		238		229	
H (1989/1996)	0.79	0.73	0.79	0.73	0.86	

Explanatory note: $H = 0$ and $H = 1$ rejected for all equations (level of confidence 99.9%).

Table A.13
Empirical results for Spain

	Standard				Weighted	
	Coeff.	t-values	Coeff.	t-values	Coeff.	t-values
INTE	0.49	*25.4*	0.46	*27.0*	0.69	*35.6*
PE	0.14	*8.2*	0.21	*15.0*	0.21	*10.3*
CE	0.11	*5.4*	0.07	*3.3*	0.01	*0.7*
TIME	2.39	*8.1*	2.84	*9.7*	2.54	9.6
LO	0.09	*13.6*			0.05	4.1
BDD	-0.02	*4.2*			-0.09	7.7
CL					-0.04	3.5
TA	-0.02	*4.5*				
Intercept	0.78	*12.2*	1.12	*23.1*	0.91	*13.0*
Adj. R-squared	0.782		0.676		0.999	
Number of observations	387		603		357	
H (1989/1996)	0.60	0.55	0.58	0.52	0.83	0.78

Explanatory note: $H = 0$ and $H = 1$ rejected for all equations (level of confidence 99.9%).

Table A.14
Empirical results for Sweden

	Standard				Weighted	
	Coeff.	t-values	Coeff.	t-values	Coeff.	t-values
INTE	0.47	*15.8*			0.71	*14.8*
PE	0.08	*6.3*			0.05	*3.4*
CE	0.04	*2.2*			0.04	*2.9*
TIME					2.04	*1.8*
LO					0.19	*4.1*
Intercept	1.55	*11.5*			0.54	*4.8*
Adj. R-squared	0.654				0.997	
Number of observations	140				137	
H (1989/1996)	0.59				0.78	0.74

Explanatory note: $H = 0$ and $H = 1$ rejected for all equations (level of confidence 99.9%).

Table A.15

Empirical results for the UK

	Standard				Weighted	
	Coeff.	*t*-values	Coeff.	*t*-values	Coeff.	*t*-values
INTE	0.63	*21.9*	0.69	*18.2*	0.65	*43.7*
PE	0.04	*2.5*	0.14	*7.6*	0.07	*7.8*
CE	0.05	*2.4*	0.01	*0.3*	0.10	*5.1*
TIME					2.94	*7.1*
LO	0.25	*18.6*			0.24	*16.8*
CL					-0.04	*3.1*
BDD					-0.05	*6.5*
TA	-0.03	*4.5*			-0.00	*2.3*
Intercept	-0.28	*2.2*	0.40	*2.7*	0.07	*0.8*
Adj. R-squared	0.706		0.475		0.999	
Number of observations	446		447		334	
H (1989/1996)	0.72		0.83		0.76	0.70

Explanatory note: $H = 0$ and $H = 1$ are rejected for all equations (level of confidence 99.9%).

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Summary

Competition and Concentration in the EU Banking Industry

This paper presents empirical evidence on the competitive structure in the banking industry in the EU as a whole as well as in individual EU countries. The study is based on the Panzar-Rosse methodology, which uses a non-structural estimation technique to evaluate the elasticity of total interest revenues with respect to changes in banks' input prices. The significant positive values of the competitiveness measure indicate that banks do not exhibit monopoly behaviour in any of the EU countries. Rather, European banking sectors operate under conditions of monopolistic competition, albeit to varying degrees. The results also provide some support for the conventional view that concentration impairs competitiveness. (JEL D41 - 43, F36, G15, G18, G21)

Zusammenfassung

Wettbewerb und Konzentration in der Kreditwirtschaft der EU

Dieser Beitrag stellt eine empirische Untersuchung der Wettbewerbsstruktur der Kreditwirtschaft in der EU insgesamt sowie in einzelnen EU-Ländern dar. Die Studie folgt der Methodik von Panzar-Rosse, bei der zur Bewertung der Elastizität der Gesamtzinseinnahmen hinsichtlich Veränderungen der Inputpreise der Banken eine nichtschematische Schätztechnik zum Einsatz gelangt. Die signifikanten positiven Meßwerte der Wettbewerbsfähigkeit zeigen, daß die Banken in keinem der EU-Länder ein Monopol-Verhalten an den Tag legen. Der Bankensektor in den europäischen Ländern arbeitet eher unter monopolistischen Wettbewerbsbedingungen, wenngleich dies von Land zu Land unterschiedlich ausgeprägt ist. Die Ergebnisse stützen bis zu einem gewissen Grade auch die konventionelle Meinung, daß Konzentration für die Wettbewerbsfähigkeit schädlich ist.

Résumé

Concurrence et concentration dans le secteur bancaire de l'UE

Cet article présente l'évidence empirique de la structure concurrentielle dans le secteur bancaire de l'ensemble de l'UE ainsi que dans chacun des pays-membres. L'étude est basée sur la méthodologie de Panzar-Rosse qui utilise une technique d'estimation non structurelle pour évaluer l'élasticité des revenus totaux d'intérêt par rapport aux variations des prix des inputs bancaires. Les valeurs positives significatives de mesure de compétitivité indiquent que les banques ne se comportent de manière monopolistique dans aucun des pays de l'UE. Plus exactement, les secteurs bancaires européens opèrent sous des conditions de concurrence monopolistique, encore qu'à des degrés différents. Les résultats soutiennent également la vue conventionnelle que la concentration affaiblit la concurrence.